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What does export diversification do for growth? An econometric Analysis

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It is frequently suggested that export diversification contributes to an acceleration of growth in developing countries. Horizontal export diversification into completely new export sectors may generate positive externalities on the rest of the economy as export oriented sectors gain from dynamic learning activities due to contacts to foreign purchasers and exposure to international competition. Vertical diversification out of primary into manufactured exports is also associated with growth since primary export sectors prevalently do not exhibit strong spillovers. Yet there have been remarkably few empirical investigations into the link between export diversification and growth. This paper attempts to examine the hypothesis that export diversification is linked to economic growth via externalities of learning-by-doing and learning-by-exporting fostered by competition in world markets. The diversification-led growth hypothesis is tested by estimating an augmented Cobb-Douglas production function on the basis of annual time series data from Chile. Based on the theory of cointegration three types of statistical methodologies are used: the Johansen trace-test, a multivariate error-correction model and the dynamic OLS procedure. Given structural changes in the Chilean economy, time series techniques considering structural breaks are applied. The estimation results suggest that export diversification plays an important role in economic growth.

I. INTRODUCTION

The idea that export diversification contributes to an acceleration of growth in developing countries is a recurrent idea in development economics. In theory, there are a number of channels through which export diversification might positively affect output growth. By increasing the number of export sectors, horizontal export diversification can reduce the dependence on a limited number of commodities that are subject to extreme price and volume fluctuations. Such swings in foreign exchange revenues may hamper efforts at economic planning, reduce import capacity, and contribute to an undersupply of investment by risk adverse producers (Dawe, 1996). Thus, decreasing export instability through horizontal export diversification may provide significant development benefits.[1] According to the Prebish-Singer thesis, vertical export diversification into manufactures may be useful if there is a general trend toward declining terms of trade for primary products (Athukorola, 2000). These arguments in favour of export diversification on the grounds that diversifying the export portfolio reduces export earnings variability and leads to terms of trade gains are based on neoclassical trade theory, which is not strictly relevant to long-run economic growth. However, it can be hypothesised that export diversification affects long-run growth as suggested by endogenous growth theory, which emphasises the role of increasing returns to scale and dynamic spillover effects (Amin Gutiérrez de Piñeres and Ferrantino, 2000).

Improved production techniques associated with export diversification are likely to benefit other industries through knowledge spillovers (Al-Marhubi, 2000).[2] The possible sources of these knowledge externalities include productivity enhancements resulting from increased competitiveness, more efficient management styles, better forms of organisation, labour training, and knowledge about technology and international markets. As Chuang (1998) argues, entering competitive international markets requires knowledge about foreign buyer's specifications, quality and delivery conditions. To satisfy these requirements, foreign purchasers help and teach local exporters to establish each stage of the production process and improve management and marketing practices. The development of efficient quality control procedures, management and marketing methods, product specifications and production guidelines is simultaneously fostered by the increased competitive pressure in world markets. If knowledge is generated through a systematic learning process initiated by exporting activities, developing countries will gain from orienting their sectors towards exporting. Hence, horizontal export diversification will have a positive net effect on aggregate output. Since manufactured exports tend to offer greater potential for sustained learning and more spillover benefits to other activities, many endogenous growth models suggest vertical diversification out of traditional primary exports into dynamic manufactured exports (Matsuyama, 1992). Accordingly, horizontal and vertical export diversification may positively affect growth.

Despite the popularity of the hypothesis of diversification-led growth there have been remarkably few empirical investigations into the implied links between export diversification and growth. To our knowledge, only Balaguer and Cantavella-Jordá (2004), De Ferranti et al. (2002), Amin Gutiérrez de Piñeres and Ferrantino (2000) and Al-Marhubi (2000) have examined the impact of export diversification on economic growth.[3] Cross-sectional studies by De Ferranti et al. (2002) and Al-Marhubi (2000) find evidence in favour of diversification-led growth. Similarly, Amin Gutiérrez de Piñeres and Ferrantino (2000: Chapter 7) find a positive link between export diversification and per capita income on the basis of panel data for Latin America countries. In their time-series analysis of structural change in exports and economic growth in Spain, Balaguer and Cantavella-Jordá (2004) also establish a positive relationship using cointegration and causality tests.[4] In contrast, the time-series studies by Amin Gutiérrez de Piñeres and Ferrantino (2000: Chapter 4, 5) show no evidence in support of diversification-induced growth in Columbia and Chile. In the case of Chile export diversification actually seems to be negatively correlated with growth.[5] However, the studies by Amin Gutiérrez de Piñeres and Ferrantino (2000: Chapter 4, 5) suffer from several methodological shortcomings.

An important problem is that the issue of cointegration, which is significant in the predicted long-run relationship between export diversification and economic growth, is not considered by Amin Gutiérrez de Piñeres and Ferrantino. The authors deal with the problem of nonstationarity of their underlying time-series by taking first differences. But if the variables of interest are cointegrated, the standard practice of taking first differences may lead to erroneous results. Another shortcoming is that Amin Gutiérrez de Piñeres and Ferrantino do not consider the presence of possible structural breaks when testing for unit roots. Neglecting structural breaks may lead to spurious unit roots. This casts some doubt on the observed unit root behaviour of the underlying series and makes their regression results additionally questionable. Finally, Amin Gutiérrez de Piñeres and Ferrantino do not conduct standard residual test for the estimated models. Without assessing the residuals for normality, autocorrelation and heteroscedasticity their regression results are little convincing.

This paper carefully investigates the long-run relationship between export diversification and growth. It attempts to test the hypothesis that export diversification is linked to economic growth via externalities of learning initiated by export activities. The study is different from the studies outlined above in several respects: First, we apply time series techniques, since evidence of significant parametric variations across countries suggest that aggregate cross country analyses may be highly misleading. Second, because standard unit root tests may be biased in the face of structural breaks, we use advanced statistical procedures that explicitly allow for structural breaks. Third, the study uses cointegration techniques to examine the long-run impact of export diversification on economic growth. Fourth, in this paper we check for the robustness of the results by utilising two different methods to estimate the parameters of the long-run relation. Given potential problems of endogeneity of the explanatory variables, one approach taken in this article considers all the included variables as potentially endogenous.

In order to investigate the diversification-led growth hypothesis we use Chilean time series data from 1962 - 2001. Chile is chosen as a case study because Chile has diversified its exports horizontally and vertically on the basis of natural resources. Since the comparative advantage of many developing countries lies in the production of resource based products, the Chile experience might demonstrate for other developing countries, if and how diversifying on the base of natural resources can accelerate their growth.

The rest of the paper is organised as follows. Section II presents a brief review of the development of the Chilean economy and of the role of export diversification in that development. In Section III the empirical model of diversification-led growth is outlined. The data and the econometric methodology are described in Section VI. The estimation results are presented in Section V. A final Section summarises the conclusions.

II. ECONOMIC DEVELOPMENT AND EXPORT DIVERSIFICATION IN CHILE

Useful and detailed surveys of the Chilean growth process are provided, among others, by Edwards and Edwards (1987) and in the book edited by Bosworth, Dornbusch and Labán (1993). In the following we present some stylised facts. First, we can observe a pattern of relatively high long-run growth, which, however, was interrupted by three deep economic crises. Chile grew by about 4.5 percent per year during the period 1963-1971, 6.8 percent from 1976-1981 and 6.1 percent on average in 1984-2001 (Figure 1). This growth performance of the Chilean economy was broken (i) by the collapse of the socialist government under President Salvador Allende ended by the military coup of 1973; (ii) by the dramatic slowdown in 1975 due to a very restrictive fiscal and monetary policy and the world economic recession; and (iii) by the deep economic depression in 1982-1983 which was associated with the general debt crisis in Latin America. The huge increase in international interest rates induced by the very tight monetary policy in the United States had devastating effects on the Chilean economy. Besides, policy mistakes such as inadequate banking supervision and a misguided effort to control inflation via the exchange rate exacerbated the recession. After the debt crises Chile started a long period of economic growth briefly interrupted by the Asian financial crisis, which hit the Chilean economy in 1998-1999.

[Figure 1]

The Chilean long-run growth performance described above was led by an increase and diversification of exports, as several authors have argued (Agosin, 1999; Ffrench Davis, 2002). In 1963-1970 exports still grew moderately by 3.6 percent. In that period Chile pursued a strategy of import substitution with few efforts to liberalise trade. However, the newly elected government under President Allende deepened the inward oriented policy, as of 1970. Under his administration (1970-1973) the Chilean Economy could be characterised as a closed economy with high import barriers and strong discrimination against exports. Export growth rates fell from 2.1 percent in 1970 to -15.1 percent in 1972 (Figure 1). After the military coup of September 11, 1973 the degree of openness of the Chilean Economy increased significantly, which was due to radical trade policy reforms implemented by the military administration under General Augusto Pinochet (1973-1989). Since 1974 exports grew very rapidly. In the seven years from 1974 to 1980, the annual growth rate of exports was 17.8 percent. Nontraditional exports also expanded, particularly those of fresh fruit, roundwood and sawnwood, and semi-manufactured copper. However, the export growth rate became negative in the period 1981-1985, with an average annual decrease of 1.5 percent, due to the appreciation of the real exchange rate and the slow down of the world economy. The second phase of high export growth rates began in 1985 after the real exchange rate had been sharply devaluated. Exports grew at an average rate 10 percent per year between 1985 and 2001. Nontraditional exports increased again as of 1985, led by agricultural products such as fresh fruits and vegetables and several wood products. Fast growing nontraditional exports also included industrial sectors, producing chemicals and basic metals machinery. Looking at the export structure over time, one can find that the degree of vertical export diversification in Chile increased sharply from about 1974 onwards. The share of manufacturing exports rose from 7 percent of the total in 1973 to 47 percent in 2001, whereas the share of copper in total exports decreased from 63 percent in 1973 to about 30 percent in 2001. But the main manufacturing exports are some few resource-based products with a low level of technological content: food products and feedstock, wood pulp and paper, and forestry products.[6] Accordingly, vertical export diversification in Chile mainly reflects the rapid expansion of certain industries. The industrial export volumes are still concentrated in few large export sectors. However, there has been a significant horizontal export diversification towards other products and more sectors. The number of products exported increased from 1440 in 1987 to 3749 in 2001 and the number of exporters rose from 3666 in 1987 to 6009 in 2001. This increase has been accompanied by a significant expansion in the number of export sectors, which rose from 91 in 1973 to 174 in 2001.[7] Although exports have become more diversified in terms of exporting sectors, most export sectors rely on natural resources. Nevertheless, several authors have suggested that there were strong knowledge spillovers from the export sectors to the rest of the economy, that fostered growth and competitiveness of other industries (Fischer, 2001; De Ferranti et al., 2002).

[Table 1]

III. EMPIRICAL MODEL

Against the background of the previous discussion of growth and export diversification in Chile, we now set out a model to test the hypothesis that export diversification is linked to economic growth via externalities of learning-by-exporting and learning-by-doing. For this purpose, we consider an economy with n sectors and $Z \in n$ export sectors. We assume there is one firm in each sector and the production of each sector $f \in [1, n]$ at any point of time t is characterised by a neoclassical production function:

$$Y_{f_t} = F_{f_t}(K_{f_t}, L_{f_t}, W_t), \qquad (1)$$

where Y_{ft} is the output of the sector. K_{ft} and L_{ft} are the conventional inputs capital and labour. W_t is the index of public knowledge which enters the production function of each sector f as a positive externality. The knowledge externality W_t has the following properties:

First, knowledge is mainly generated within the export sectors of the economy as a result of learning-by-exporting and learning-by-doing activities. The idea behind learning-by-exporting is that exporters gain from the knowledge base of their buyers as foreign purchasers offer advice on productivity enhancements. Learning-by-doing is associated with knowledge creation as a side product of production, depending on the firm's cumulative output. Thus, an export-induced expansion of the firm's output increases its stock of knowledge. The process of knowledge generation is simultaneously accelerated by the competitive pressure of the international marketplace. For simplicity it is assumed that each one of the Z_t export sectors produces the same amount of knowledge W_e, so that the level of aggregate knowledge can be written as

$$W_t = Z_t W_{e_t}. \tag{2}$$

Due to the fact that W_{et} is taken as a constant parameter,[8] the level of knowledge in the economy can be expressed as a function of the number of export sectors without including W_{et}:

$$W_t = G(Z)_t. \tag{3}$$

However, as many authors have argued, learning effects may depend on the structure of exports.[9] In particular, since primary exports may not have a high potential for learning-by-doing and learning-by-exporting, knowledge creation is expected to increase with the share of manufactured products in total exports. Hence, the knowledge externality that we consider here takes the form

$$W_t = G(Z_t, IX_t), \tag{4}$$

where the share of manufactured exports in total exports (IX_t) and the number of export sectors (Z_t) are proxies for the stock of knowledge in the economy.

Second, knowledge, W_t, is a public good that is regarded as constant within all sectors. We assume that W_t affects all sectors equally but how W_t affects function F_f is neglected by the export sectors. Treating W_t as given, F_f behaves like a constant-returns-to-scale production function. Let there be perfect competition in the sense that all firms are price takers, and set

$$Y_t = \sum_{f=1}^{n} Y_{f\,t}, \quad K_t = \sum_{f=1}^{n} K_{f\,t}, \quad L_t = \sum_{f=1}^{n} L_{f\,t}, \tag{5}$$

the total production Y_t in the economy can be written as

$$Y_t = \sum_{f=1}^{n} Y_{f\,t} = F_t(K_t, L_t, W_t) = F_t(K_t, L_t)G(Z_t, IX_t) = K_t^{\alpha} L_t^{\beta} Z_t^{\gamma} IX^{\delta}, \tag{6}$$

where K_t represents the stock of accumulated capital, L_t is the labour force of the economy and the parameters α, β, γ, δ are constants. Adding the number of export sectors and the share of manufactured exports in total exports as explanatory variables in equation (6) implies that horizontal and vertical export diversification are linked to economic growth via externalities of learning-by-doing and learning-by-exporting (since γ, $\delta > 0$). To investigate the long-run relationship between export diversification and economic growth along with capital and labour, equation (6) is expressed in the following log-linear regression form:

$$LY_t = c + \alpha LK_t + \beta LL_t + \gamma LZ_t + \delta LIX_t + e_t, \tag{7}$$

where L represents the natural logarithms of the variables, and e_t is the usual error term representing variables not included in the model, exogenous shocks, and errors of measurement; e_t is assumed to be white-noise and normally and identically distributed. The log-linear specification implies that the estimates of α, β, γ and δ are elasticities according to equation (6). Therefore, a simple, testable, and theoretically consistent test for the diversification-led growth Hypothesis is:

$$H_0: \quad \gamma, \delta = 0$$
$$H_1: \quad \gamma, \delta > 0 \quad .$$

Consequently, the diversification-led growth hypothesis will not be rejected by the data if the estimates of γ and δ are positive and statistically significant.

IV. DATA AND ECONOMETRIC METHODOLOGY

The data used to estimate equation (7) are annual for the period 1962-2001 (T = 40, 1≤ t≤ 40). The aggregate output (Y_t) is measured by the Chilean GDP. The Chilean capital stock (K_t) was calculated on the basis of accumulated capital expenditure using the perpetual inventory method in simple form. GDP and capital stock are evaluated at constant prices (1996 prices). The data on labour (L_t) corresponds to the number of people employed in each year. The ratio of manufactured exports to total exports (IX_t) was computed on the basis of real industrial exports and real aggregate exports (1996 prices). Z_t is the number of export sectors classified by the *Standard International Trade Classification* at the three-digit level. With the exception of the number of export sectors, which are from the *United Nations* (COMTRADE), the data used in this study are from *Banco Central de Chile*. Figure 2 shows the evolution of the data between 1962 and 2001. All data are in logarithmic forms.

[Figure 2]

From Figure 2, it can be inferred that all series are trending and thus nonstationary. Nonstationary time series may contain unit roots. Such time series are said to be integrated of order d, $I(d>0)$, because they have to be differenced d times to achieve stationarity (difference stationary series). In the case where nonstationary time series are not driven by a unit root process, they are subject to deterministic time trends (trend stationary series). By removing the deterministic trend they can be made stationary, $I(0)$. The trending behaviour of the underlying series is investigated by means of unit root tests. However, standard unit root tests, such as the Augmented-Dickey-Fuller test, may be biased in favour of identifying data as integrated if there are structural changes (Perron, 1989). For all the series there is indeed a strong likelihood that structural discontinuities are present (e.g. the socialist government of President Allende (1970-1973), the 1975 recession, and the Latin American depth crises (1981)). Therefore, we undertake the unit root test developed by Perron (1997). The Perron procedure permits a formal evaluation of the time series properties in the presence of structural breaks at unknown points in time. It allows the break date to be identified endogenously by the testing procedure itself. However, the Perron procedure allows only for one possible break point for any single series. To consider the possibility that two break points occurred over the relevant period we apply Kapetanios' (2002) test for the unit root hypothesis against the alternative of trend stationarity with two endogenously determined breaks.

By means of these test procedures LY_t, LK_t, and LL_t are found to be $I(1)$ variables whereas LZ_t and LIX_t are stationary around a deterministic trend. The trend stationary series are then transformed into stationary series, $I(0)$, by extracting the trend. To test for the existence of a long-run relationship among LY_t, LK_t, and LL_t the multivariate cointegration technique developed by Johansen (1995) is employed. As each $I(0)$ variable creates an additional cointegration vector, the $I(0)$ variables are separated from the $I(1)$ variables in testing for cointegration rank. After having established the existence of a long run relationship between LY_t, LK_t, and LL_t we include the $I(0)$ variables in the long-run relationship. Following Lütkepohl and Wolters (1998) we use the error correction formulation first outlined by Stock (1987) to estimate a long-run relationship among $I(1)$ and $I(0)$ variables. However, the endogeneity of any of the regressors may influence, asymptotically, the robustness of the estimates. Since some of the variables are potentially endogenous, the dynamic ordinary least squares (DOLS) proposed by Saikkonen (1991) and Stock and Watson (1993) is applied. The DOLS procedure has the advantage to provide unbiased and asymptotically efficient estimates of long run relations, such as equation (7), even in the presence of endogenous regressors.

V. EMPIRICAL ANALYSIS

Time series properties

The determination of the order integration of LY_t, LK_t, LL_t, LZ_t, and LIX_t is crucial when carrying out the analysis by means of the Johansen, the ECM and the DOLS procedures. It is well known that standard unit root tests are not be able to reject the unit root hypothesis if the deterministic trend of a series has a break.[10] The methodology developed by Perron (1997) can distinguish the unit root hypothesis from that of a trend-stationary series with a single break. In order to test the unit root null hypothesis against the one-break alternative, we estimate two models of the Dickey-Fuller type without any prior knowledge of any potential break dates, i.e.

$$y_{1_t} = \mu_1 + \theta_1 DU_t + b_1 t + \delta_1 D(TB)_t + a_1 y_{t-1} + \sum_{i=1}^{k} c_{1_i} \Delta y_{t-1} + e_{1_t}, \tag{8}$$

$$y_{2_t} = \mu_2 + b_2 t + \delta_2 DT_t + \hat{y}_t, \tag{9a}$$

$$\hat{y}_t = a_2 \hat{y}_{t-1} + \sum_{i=1}^{k} c_{2_i} \Delta \hat{y}_{t-1} + e_{2_t}, \tag{9b}$$

where y_{1t} and y_{2t} are the series of interest, Δ is a difference operator, $TB \in T$ denotes the time at which the change in the trend function occurs and DU_t = 1(t>TB), $D(TB)_t$

$=1(t=TB+1)$, $DT_t = 1(t>\mathrm{TB})(t\text{-}TB)$ are indicator dummy variables for the break at time TB. The regression models (8) and (9) correspond, respectively to the crash model and the changing growth model proposed by Perron (1989). Model (8), the *innovational outlier model*, allows for a one-time change in the intercept of the trend function. It involves a one step regression by estimating the trend function and the dynamics of the process simultaneously. Model (9), the *additive outlier model*, which involves a two step regression, allows for a change in the slope of the trend function without a change in the level.[11] For LY_t, LL_t LZ_t, and LIX_t regression of type (8) is carried out. Regression (9) is applied to LK_t as the capital stock data indicates no "crash" but a change in the slope of the series.

The break point is chosen by estimating the models for each possible break date in the data set and TB is selected as the value which minimises the t-statistics for testing $a = 1$. Accordingly, the estimated break point TB^* corresponds to the date for which the t-statistic is minimised under the unit root hypothesis: $t^*_a(i) = \mathrm{Min}_{TB}\, t_{\hat{a}}(i, TB, k)$, where $t_{\hat{a}}(i, TB, k)$ is the t-statistic for testing $a = 1$ under model $i = 1,2$ with a break date TB and truncation lag parameter k. If $\mathrm{Min}_{TB}\, t_{\hat{a}}(i, TB, k)$ exceeds (in absolute value) its critical value reported by Perron (1997), the hypothesis of difference stationarity and a unit root is rejected.

Since considerable evidence exists that data-dependent methods to select the value of the truncation lag k are superior to choosing a fixed k a priori, we follow Perron (1997) and use the t-sig method. Here, k max is specified to be four. If the last included lag is insignificant, the number of lags is reduced by one and the equation is reestimated, until a significant lagged dependent variable is found. If none of the coefficients on the lagged variables are found to be significant (at the 10% level), no lags are utilised in the test. Table 2 contains the results of the sequential unit root tests for the variables in levels and in first differences.[12] The results indicate that LY_t, LK_t, and LL_t are integrated of order one, whereas the export sector and the industrial export share series (LZ_t, LIX_t) are trend stationary with at least one structural break in 1972.

[Table 2]

However, we do need to be cautious in interpreting the results. As Lumsdaine and Papell (1997) point out, results regarding tests of the unit root hypothesis are sensitive to the assumed structural breaks. The authors show that the results obtained using one endogenous break are often reversed when a model with two breaks is estimated. This introduces a degree of uncertainty in the analysis. Therefore we check the validity of the results represented in Table 2 by considering the possibility that two break points occurred over the relevant time period. We employ Kapetanios' (2002) test for the null hypothesis of a unit root against the alternative hypothesis of an unspecified number of structural breaks. We estimate two models:

$$y_{1_t} = \mu_1 + b_1 t + a_1 y_{t-1} + \sum_{i=1}^{m} \delta_1 DU_{i,t} + \sum_{i=1}^{k} c_{1_i} \Delta y_{t-1} + e_{1_t}, \tag{10}$$

$$y_{2_t} = \mu_2 + b_2 t + a_2 y_{t-1} + \sum_{i=1}^{m} \delta_2 DT_{i,t} + \sum_{i=1}^{k} c_{2_i} \Delta y_{t-1} + e_{2_t}, \tag{11}$$

where y_t is the variable considered, m denotes the number of breaks, and $DU_{i,t}$ and $DT_{i,t}$ are defined as in equation (8) and (9). Setting $m = 2$, model (10) allows for two breaks in the intercept of the trend function. In model (11) the two breaks are restricted to the slope of the trend function. Since visual inspection of the capital stock data suggests only possible changes in the slope, regression (11) is applied to LK_t. For LY_t, LL_t LZ_t, and LIX_t we carry out regression of type (10), where both breaks in the trend function are restricted to the intercept. Running the regressions for all indicator dummy variables, we chose the date of the first structural break such that the sum of squared residuals is smallest among all possible break points in the data set. Imposing the estimated break date on the sample, we start looking for the second break. Again, the second break point is associated with the minimum of squared residuals.

The results of testing the unit root null against the two-break alternative are reported in Table 3. Except for the selected break points, they do not differ from the results of the Perron (1997) procedure. As it is more plausible that there are two breaks in the export diversification data (the first for the strong discrimination against exports as of 1971 and the second before the rapid trade liberalisation and the spectacular increase in diversification of exports during 1974) we summarise the main points of our results as follows:

[Table 3]

The results show clearly that there is a mixture of $I(1)$ and $I(0)$ variables. The null hypothesis of a unit root cannot be rejected for LY_t, LK_t, and LL_t in levels. Since for the first differences the unit root hypothesis can be rejected, it is concluded that real GDP, aggregate capital and employed people are integrated of order one, $I(1)$. For LIX_t and LZ_t, however, there is strong empirical evidence that these variables are trend stationary, interrupted by trend breaks in 1971 and 1973. The unit root hypothesis can be rejected in favour of broken trend stationarity at the 1% significance level.[13]

Since trend stationarity may exacerbate potential problems of multicollinearity between LIX_t and LZ_t, we use the detrended data, denoted in the following as lix_t and lz_t.[14] That is to say, to estimate the parameters α, β, γ and δ of equation (7) we take the residuals (lix_t and lz_t) from a regression of LIX_t and LZ_t on a constant, a time trend and two indicator dummy variables ($DU_{t,i}$) for structural breaks in 1971 and 1973.[15] As the export diversification data are detrended they can be regarded as stationary.[16] For every stationary variable included in Johansen's test of cointegration among $I(1)$ variables the cointegration rank will increase accordingly. In order to avoid problems in identifying which of the possible cointegration vectors might present the stationary series, we separate the $I(0)$ variables from the cointegration analysis in the next section.

Testing for cointegration

Having confirmed the existence of a unit root in the GDP, the capital stock, and the labour series, the multivariate cointegration technique developed by Johansen (1995) is applied to examine the long-run relationship among LY_t, LK_t, and LL_t. The Johansen approach estimates long-run or cointegration relationships between $I(1)$ series using a maximum likelihood procedure, which tests for the number of cointegration relationships. The method is based on the unrestricted vector autoregression (VAR) model represented by the following equation:

$$y_t = \mu + \sum_{k=1}^{p} \Pi_k y_{t-k} + \varepsilon_t , \tag{12}$$

where y_t is an $(n \times 1)$ column vector of n $I(1)$ Variables, Π_k is a coefficient matrix, μ represents an $(1 \times n)$ vector of constants, p denotes the lag length, and ε_t is a disturbance term independently and identically distributed with zero mean and constant variance. Since y_t = [LY_t, LK_t, and LL_t]' is assumed to be $I(1)$, letting $\Delta y_t = y_t\text{-}y_{t-1}$, equation (12) can be rewritten in first difference notation reformulated in error correction form (ECM) as:

$$\Delta y_t = \mu + \sum_{k=1}^{p-1} \Gamma_k \Delta y_{t-k} + \Pi y_{t-1} + \varepsilon_t , \tag{13}$$

where Γ_k and Π represent coefficient matrices and the rank r of matrix Π determines the number of cointegration relations in the system. As Δy_t and $\Delta y_{t\text{-}1}$ variables are $I(0)$ and $y_{t\text{-}1}$ variables are $I(1)$, equation (13) will be balanced if left-hand side and right hand-side have the same degree of integration. This will either occur if $r = 0$, so that $\Pi = 0$, in which case the variables in y_t are not cointegrated or if the parameters of Π are such that $\Pi y_{t\text{-}1}$ is also $I(0)$. In the first case ($r = 0$; $\Pi = 0$) equation (13) is just a traditional VAR model in first differences. The latter case applies when the rank of Π is greater than zero, indicating that there will exist $r < n$ cointegration relations, meaning r possible stationary linear combinations of y_t. If $0 < r < n$, the reduced-rank matrix Π can be decomposed into two matrices α and β (each n × r), such that $\Pi = \alpha\beta'$. The term $\alpha\beta' y_{t\text{-}1}$ is the error correction term with $\beta' y_{t\text{-}1}$ representing the cointegration relations and α being the loading matrix α of adjustments coefficients containing the weights of the cointegration relations. The cointegrating vector β has the property that $\beta' y_t$ is stationary even though y_t itself is nonstationary. If, on the other hand, the matrix Π has a full rank $r = n$, then all n components of y_t are stationary.

The number of cointegrating vectors (the cointegration rank), r, can be formally tested with the trace statistics. The trace statistic tests the null hypothesis that the number of distinct cointegration vectors is less or equal to r against a general alternative. Asymptotic critical values for testing the null hypothesis are provided in Osterwald-Lenum (1992). The lag length p is chosen such as to minimise the Hannan-Quinn and the Schwarz criterion ($p = 2$). According to the trace test (Table 4) the hypothesis that only one cointegrating vector is in the system of $I(1)$ variables cannot be rejected at the 1% significance level. Thus, LY_t, LK_t, and LL_t are cointegrated, which implies a long-run relation between the number of employed people, capital stock and real GDP (in logarithms).

[Table 4]

Having found that the relation LY_t - LK_t - LL_t is stationary, we also determine the cointegration rank in the system of n = 2 export diversification variables, lix_t and lz_t. Results are presented in Table 5. According to the trace test there exist r = 2 cointegrating vectors, which implies a full-rank matrix Π. This result confirms our earlier conclusion that lix_t and lz_t are stationary. In the following we will include lix_t and lz_t in the long run relation between LY_t, LK_t, and LL_t by fitting an error correction model to these variables.

[Table 5]

Estimation of the long-run elasticities: Error correction model results

We employ the one step error correction model according to the technique of Stock (1987) to estimate the coefficients of the long-run relation between export diversification and economic growth along with capital and labour. The estimation is based on the Bewley (1979) transformed single equation form of equation (13). Since lix_t and lz_t are stationary and the relation LY_t - LK_t - LL_t is also stationary, it is possible to include these variables in a single equation error correction model. In our case we regress ΔLY_t on LY_{t-1}, LK_{t-1}, and LL_{t-1}, all differences of this variables up to lag order two, the detrended export diversification series (lix_t and lz_t) also up to lag order two,[17] an intercept term, a step dummy $du75$ and an impulse dummy $d75$.[18] LK_t, LL_t, lix_t, and lz_t are assumed to be weakly exogenous. The following equation results by applying Hendry's general-to-specific approach, where successively the least significant variables are eliminated until there remain only coefficients significant at the 5%-level:[19]

$$\begin{aligned}
\Delta LY_t &= \underset{(1.01)}{0.038} - \underset{(-7.27)}{0.874^{***}} LY_{t-1} + \underset{(6.53)}{0.602} LK_{t-1}{}^{***} + \underset{(4.94)}{0,499^{***}} LL_{t-1} \\
&\quad + \underset{(9,07)}{2.619^{***}} \Delta LK_t - \underset{(-3.13)}{1.303^{***}} \Delta LK_{t-2} - \underset{(-2.68)}{0.503^{**}} \Delta LL_{t-1} \\
&\quad + \underset{(3,13)}{0.387^{***}} lz_t + \underset{(2.83)}{0.097^{***}} lix_t + \underset{(3.22)}{0.083^{***}} lix_{t-2} \\
&\quad - \underset{(-3.41)}{0.071^{***}} du75 - \underset{(-4.44)}{0.114^{***}} d75
\end{aligned} \tag{14}$$

$\overline{R}^2 = 0.91$ $SE = 0.021$ $DW = 1.94$ $JB = 1.78\,(0.17)$
$ARCH(1) = 0.93\,(0.34)$ $ARCH(2) = 0.79\,(0.46)$ $ARCH(4) = 0.84\,(0.51)$
$LM(1) = 0.17\,(0.68)$ $LM(3) = 0.26\,(0.85)$ $White = 1.33\,(0.27)$

We interpret the coefficient of LY_{t-1} as significant at the 1% level, as we have already established the existence of a cointegration relationship between the number of occupied people, capital stock and real GDP (in logarithms).[20] Normalising on the coefficient of LY_{t-1} in (14) gives the following long-run relation:

$$LY_t = 0.69LK_t + 0.57LL_t\,. \tag{15}$$

Since the coefficients of lz_t, lix_t and lix_{t-2} are positive and highly significant, the diversification-led growth hypothesis can not be rejected. Adding the long-run impact of horizontal and vertical export diversification normalised on real GDP yields equation (16):

$$LY = 0.69LK + 0.57LL + 0.44Lz + 0.21Lix\,. \tag{16}$$

From equation (16), it can be inferred that Chilean GDP increases by 0.44 percent in response to a one percentage increase in numbers in export sectors. A one-percent increase in the share of manufactured products in total exports results in a 0.21 average percent increase in GDP. This indicates that (horizontal and vertical) export diversification plays an important role in economic growth. According to the estimates the contribution of capital to GDP is more significant than labour. This feature is in line with economic theory that suggests that opening to trade and the elimination of distortions increase the average quality of capital and improve the allocation of capital towards sectors with higher marginal productivity. A further reason for capital stock growth to be more important for GDP growth is that the structural base of Chile, like other developing countries, may be characterised by capital shortage and labour abundance. Any further growth in the labour force would therefore not contribute to economic growth as much as growth in the capital stock.

However, it is important to emphasise that the right hand side variables of equation (16) are assumed to be weakly exogenous. If the regressors are not weakly exogenous, the single equation ECM will be biased and inefficient and *t*-tests based on the model parameters will be highly misleading. In that, we cannot be sure that economic growth in Chile is really driven by export diversification.[21]

Estimation of the long-run elasticities: Dynamic OLS results

To check for the robustness of the estimates, we apply the Dynamic OLS (DOLS) procedure developed by Saikkonen (1991) and Stock and Watson (1993). The use of this procedure ensures that our estimates are valid even if some of the explanatory variables in (16) are endogenous. Furthermore, the procedure allows for a direct estimation of a mixture of *I*(1) and *I*(0) Variables. It is asymptotically equivalent to

Johansen's (1995) maximum likelihood estimator and is known to perform well in small samples like ours. The DOLS regression in our case is given by equation (17) below:

$$LY_t = \mu + \alpha LK_t + \beta LL_t + \gamma lz_t + \delta lix_t + \sum_{i=-2}^{i=2} \Phi_1 \Delta LK_{t+i} + \sum_{i=-2}^{i=2} \Phi_2 \Delta LL_{t+i} + du75 + d75 + \varepsilon_t, \tag{17}$$

where α, β, γ, and δ are the long-run elasticities and Φ_1, Φ_2 are coefficients of lead and lag differences of the $I(1)$ regressors, which are treated as nuisance parameters. These serve to adjust for possible endogeneity, autocorrelation, and nonnormal residuals and result in consistent estimates of α, β, γ, and δ. Similar to regression (14) the dynamic OLS is carried out up to second order of leads and lags.[22] The results of the DOLS procedure are presented in Table 6.[23] The diagnostics tests statistics underneath Table 6 do not indicate any problems with autocorrelation, heteroscedasticity or nonnormality. All p-values exceed usual (5%) significance levels.[24] Thus, valid inference can be drawn from the estimated elasticities:

[Table 6]

Again, the results in table 6 show that both vertical and horizontal export diversification significantly influence Chilean growth along with capital and labour. The estimated elasticities α, β, γ, δ are positive and statistically significant. The magnitude of the coefficients in Table 6 does not differ substantially from equation (16), except for the coefficients of *Lix*. Compared to equation (16), Table 6 contains a much lower elasticity of vertical export diversification. The most obvious explanation for the large difference between the coefficients (0.21 and 0.08) is that the share of manufactured products in exports is likely to be not weakly exogenous. If there are potential feedback relations between LY_t, LK_t, LL_t, lz_t, and lix_t, then in equation (16) the estimated contribution of vertical export diversification to GDP is biased. For that reason the elasticity of vertical export diversification in equation (16) is likely to exceed its "true" value in Table 6. Similar to equation (16), a one percentage increase in numbers in export sectors results in about 0.5 average percent increase in GDP. Thus, the relationship between the degree of horizontal export diversification and aggregate output is economically large. This finding is in line with the results of Al-Marhubi (2000), who used the same indicator to measure export diversification.[25] In connection with the theoretical foundations underpinning our model, the estimation results argue for the hypothesis that horizontal export diversification is linked to economic growth via externalities of learning activities. These learning activities lead to improved production techniques, more efficient management styles, and better forms of organisation benefiting the economy as a whole. Interestingly, the estimated elasticity of horizontal export diversification (γ = 0.49) is much higher than the elasticity of vertical export diversification (δ = 0.08). Accordingly, orienting sectors towards exporting is more important for economic growth than rising the share of manufacturing exports in total exports. Though we should emphasise that vertical export diversification in Chile mainly reflects the rapid expansion of some few resource-based industries with low or medium levels of technology such as food and feedstock, wood and forestry products. Further diversification of Chile's exports towards a wide range of manufacturing products with higher technological contents possibly generates stronger growth effects. Nevertheless, the Chilean case demonstrates that export diversification on the basis of natural resources can accelerate growth, since most of Chile's export sectors rely on natural resources.

VI. SUMMARY AND CONCLUSIONS

In this study the diversification-led growth hypothesis is tested by estimating an augmented Cobb-Douglas production function on the basis of time series data from Chile. Statistical procedures are used to test for a unit root in the underlying series by considering the possibility that structural breaks at unknown time points occurred over the period 1962-2001. The results indicate that all but two series are integrated. To test for cointegration between the integrated series of order one, the multivariate cointegration methodology proposed by Johansen is used. Having established cointegration between the $I(1)$-variables, an error correction model is fitted to the series of different order of integration to estimate the long-run relationship between export diversification and economic growth. To check for the robustness of the estimate the DOLS procedure is applied. In contrast to existing time-series studies, the estimates suggest that export diversification plays an important role in economic growth. This result is robust to different estimation techniques and is in conformity with the hypothesis that export diversification is linked to economic growth via externalities of learning activities set off by exporting. An interesting finding is that orienting further sectors towards exporting is more important for growth than increasing the share of industrial exports in total exports. However, this finding must be regarded against the concrete background of vertical export diversification in Chile. Vertical export diversification in Chile mainly reflects the rapid expansion of certain resource-based industries in particular those that export food products and feedstock. Therefore, industrial export volumes are still concentrated in few large resource-based sectors with low or medium levels of technology. Diversifying and increasing industrial exports with higher technological contents possibly generates stronger growth effects. Nevertheless, a noteworthy conclusion of this paper is that export diversification on the basis of natural resources can play an important role in the growth process of developing countries, which are dependent on agricultural and mining exports. Since most of Chilean export sectors rely on natural resources, lessons for other developing countries can be drawn from the Chilean experience with regard to resource-based diversification strategies. For Chile itself, there exists the danger that the resource-based export diversification gradually wears out. Efforts should be made to establish nonresource-based sectors with higher technological opportunities in order to sustain the process of export diversification and economic growth.

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NOTES

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[1] The link between export diversification and export earnings instability has been the subject of considerable research in the last two decades. See Stanley and Bunnag (2001) for a review of the theoretical and empirical literature on this topic.

[2] See Amin Gutiérrez de Piñeres and Ferrantino (2000: Chapter 8) for an endogenous growth model, in which technological or marketing knowledge in one export sector diffuses into other lines of exporting.

[3] The authors use several indicators for export diversification, such as, for example, the number of export sectors or the Herfindahl index.

[4] Balaguer and Cantavella-Jordá (2004) consider the impact of structural transformation from traditional primary exports to nontraditional manufactured exports on Spanish GDP and thus the impact of vertical export diversification on growth.

[5] Amin Gutiérrez de Piñeres and Ferrantino (2000: Chapter 4) use the Herfindahl index to measure export concentration. The correlation between export concentration and Chilean output turns out to be statistically significant. The coefficient of the Herfindahl index has not the expected negative sign but is positive, which implies a negative correlation between export diversification and aggregate output.

[6] The Central Bank of Chile classifies the Chilean manufacturing exports according to the comprehensive definition of manufacturing of the ISIC.

[7] The declaration refers to three digit export sectors according to the SITC definition.

[8] It is empirically not directly observable.

[9] See, for example, Chuang (1998), Matsuyama (1991).

[10] Augmented Dickey-Fuller and Phillips-Perron tests would indicate that each series is integrated of order one. (Results are not reported here). However, the observed unit root behaviour is the result of failure to account for structural changes.

[11] The *additive outlier model* implies that the change in the trend function is sudden. The *innovational outlier model* implies that the break in the series does occur gradually.

[12] All our empirical tests have been carried out by EVIEWS 5.0.

[13] The unit root tests proposed Lumsdaine and Papell (1997) also indicate that real GDP, aggregate capital and occupied people are integrated of order one, whereas the export sector and the industrial share series can be constructed as stationary fluctuations around a breaking trend function. Like above, the selected breaks years in the export sector and the industrial share series are 1971 and 1973. The details of the tests are not reported for brevity, but are available upon request.

[14] Collinearity between lz_t and lix_t was investigated by inspecting the correlation matrix. The correlation coefficient of 0.50 indicates a low degree of collinearity between the detrended series. In contrast, if we compute the correlation matrix of the trended series (LZ_t and LIX_t) we have a correlation coefficient of 0.96, indicating a very high degree of collinearity.

[15] To assess the structural stability of the trend stationary models, we additionally calculated the recursive residuals. Recursive residual analysis also suggests that there are structural breaks in 1971 and 1973.

[16] Results in the next section further confirm that lz_t and lix_t can be regarded as stationary.

[17] The lag length was determined using the Hannan-Quinn and the Schwarz criterion.

[18] $du75$ is 1 from 1975 onwards and zero before 1975; $d75$ is 1 in 1975. The possible reason for $du75$ and $d75$ to be important is the deep economic depression in 1975.

[19] t-ratios in parentheses underneath the estimated coefficients. ** and*** denote the 5% and 1% level of significance respectively. The number in parenthesis behind the values of the diagnostic tests statistics are the corresponding p-values. *JB* is the Jarque-Bera test for normality, *LM* (k), k=1,3, are *LM* tests for autocorrelation based on 1 and 3 lags, respectively and *ARCH* (k) is an LM test for autoregressive conditional heteroscedasticity of order k =1, 2, 4. *White* = White test for heteroscedasticity of the errors.

[20] Conventional distributional results are applicable for the t-test statistic since the Bewley-transformed ECM term is stationary (according to the trace test). Additionally, one may argue that the null of no cointegration may be rejected at the 1% significance level, because the t-value of the loading coefficient (-7.27) lies below the critical value for two stochastic regressors (-4.38) according to the test for cointegration suggested *inter alia* by Ericcson and MacKinnon (2002). However, further stationary variables may influence the distribution of the ECM test statistic under the null of no cointegration.

[21] Tests for weak exogeneity within the Johansen framework indicate that LK_t is weakly exogenous, while LY_t and LL_t are endogenous. However this test is not invariant to the inclusion of stationary variables, such as Lz_t, Lix_t. Thus, weak exogeneity in the full system (LY_t, LK_t, LL_t, Lz_t, Lix_t) may differ from weak exogeneity in the subsystem (LY_t, LK_t, LL_t). Instead of investigating the weak exogeneity status of each of the "explanatory" variables, the DOLS procedure is preferred here.

[22] Dummy variables are used to capture the effects of the deep economic crises in 1975; *du*75 and *d*75 are defined as in equation (14).

[23] t-ratios in parentheses underneath the estimated coefficients. ** and*** denote the 5% and 1% level of significance respectively. The number in parenthesis behind the values of the diagnostic tests statistics are the corresponding p-values. *JB* is the Jarque-Bera test for normality, *LM* (k), k=1,3, are *LM* tests for autocorrelation based on 1 and 3 lags, respectively and *ARCH* (k) is an LM test for autoregressive conditional heteroscedasticity of order k =1, 2, 4. *White* = White test for heteroscedasticity of the errors.

[24] Following Stock and Watson (1993) the insignificant leads and lags were not dropped. If we follow Hendry`s general-to-specific approach the residuals appear not to be as free of autoregressive conditional heteroscedasticity, although the coefficients for the explanatory variables are reasonably similar.

[25] However, the results are not directly comparable due to different estimation methods and different economic variables in the estimation equations.

REFERENCES

Agosin, M. (1999) Trade and growth in Chile, *Cepal Review*, **68**, 79-100.

Al-Marhubi, F. (2000) Export diversification and growth: An empirical investigation, *Applied Economics Letters*, **7**, 559-562.

Amin Guitiérrez de Piñeres, S. and M. J. Ferrantino (2000) *Export dynamics and economic growth in Latin America*, Ashgate Publishing Ltd, Burlington, Vermont.

Athukorola, P.C. (2000) Manufacturing exports and terms of trade of developing countries: Evidence from Sri Lanka, *Journal of Development Studies*, **36**, 89-104.

Balaguer, J. and M. Cantavella-Jordá (2004) Structural change in exports and economic growth: cointegration and causality analysis for Spain (1961-2000), *Applied Economics*, **36**, 473-477.

Banco Central de Chile (2001) *Indicadores económicos y sociales de Chile 1960-2000*, Santiago de Chile.

Banco Central de Chile (1980-2002) *Boletín mensual* (various issues), Santiago de Chile.

Bewley, R. A. (1979) The direct estimation of the equilibrium response in a linear model, *Economics Letters*, **3**, 357-361.

Bosworth, B., R. Dornbusch and R. Labán (1993) *The Chilean economy: policy lessons and challenges*, Washington D.C.: The Brookings Institution.

Chuang, Y-C. (1998) Learning by doing, the technology gap, and growth, *International Economic Review*, **39**, 697-721.

Dawe, D. (1996) A new look at the effects of export instability on investment and growth, *World Development*, **24**, 1905-1914.

De Ferranti, D., G. E. Perry, D. Lederman and W. F. Maloney (2002) *From natural resources to the knowledge economy*, The World Bank. Washington D.C.

Edwards, A. and S. Edwards (1987) *Monetarism and liberalisation: The Chilean experiment*, Chicago: The University of Chicago Press.

Erricsson, N. and J. MacKinnon (2002) Distributions of error correction test for cointegration, *Econometrics Journal*, **5**, 285-318.

Ffrench-Davis, R. (2002) El impacto de las exportaciones sobre el crecimiento en Chile, *Revista de la Cepal*, **76**, 143-160.

Fischer, R. (2001) Trade liberalisation, development and government policy in Chile, Documentos de Trabajo 102, Centro de Economía Aplicada, Universidad de Chile.

Johansen, S. (1995) *Likelihood-based inference in cointegrated vector autoregessive model*s, Oxford University Press, New York.

Kapetanios, G. (2002) Unit root testing against the alternative hypothesis of up to m structural breaks, Queen Mary University of London, Department of Economics, Working Paper No. 469.

Lumsdaine, R. L. and D. H. Papell (1997) Multiple trend breaks and the unit root hypothesis, *Review of Economics and Statistics*, **79**, 212-218.

Lütkepohl. H. and J. Wolters (1998) A money demand system for German M3, *Empirical Economics*, **23**, 371-386.

MacKinnon, J. G. (1991) Critical values for cointegration tests, in: Engle, R. F., C. W. J. Granger (eds.): *Long-run economic relationships: Readings in cointegration*, Oxford University Press, New York.

Matsuyama, K. (1992) Agricultural productivity, comparative advantage, and economic growth, *Journal of Economic Theory*, **58**, 317-334.

Osterwald-Lenum, M. (1992) A note on quantiles of the asymptotic distribution of the maximum likelihood cointegration rank test statistic, *Oxford Bulletin of Economics and Statistics*, **54**, 461-472.

Perron, P. (1989) The great crash, the oil price shock, and unit root hypothesis, *Econometrica*, **57**, 1361-1401.

Perron, P. (1997) Further evidence on breaking trend functions in macroeconomic variables, *Journal of Econometrics*, **80**, 355-385.

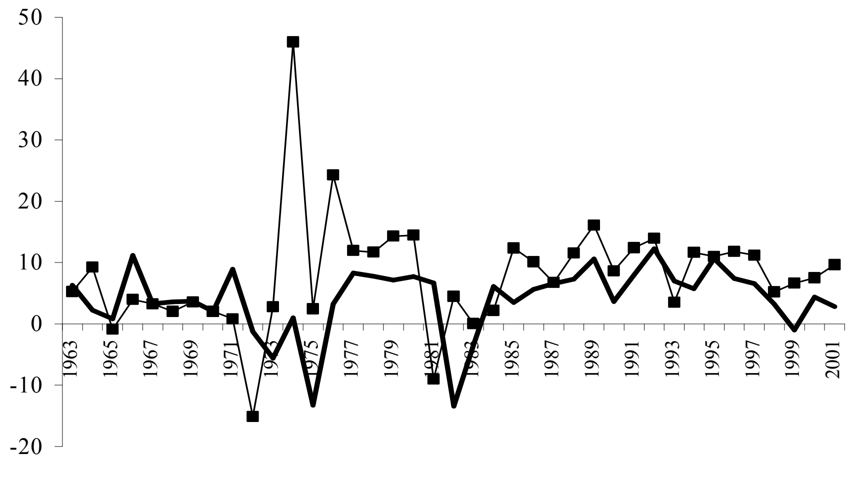
Saikkonen, P. (1991) Asymptotically efficient estimation of cointegrating regressions, *Econometric Theory*, **7**, 1-21.

Stanley, D. L. and S. Bunnag (2001): A new look at the benefits of diversification: Lessons from Central America, *Applied Economics*, **33**, 1369-1383.

Stock, J. H. (1987): Asymptotic properties of least squares estimators of cointegrating vectors, *Econometrica*, **55**, 1035-1056.

Stock, J. H., M. W. Watson (1993): A simple estimator of cointegrating vectors in higher order integrated systems, *Econometrica*, **61**, 783-820.

United Nations (2003): *Commodity Trade Statistics Database* (COMTRADE). New York.



real export growth (▪▪▪), real GDP growth (▬)

Source. Calculated on the basis of data from *Banco Central de Chile* (based on constant 1996 prices).

Fig. 1. *Export and GDP growth rates (in percent), 1963-2001*

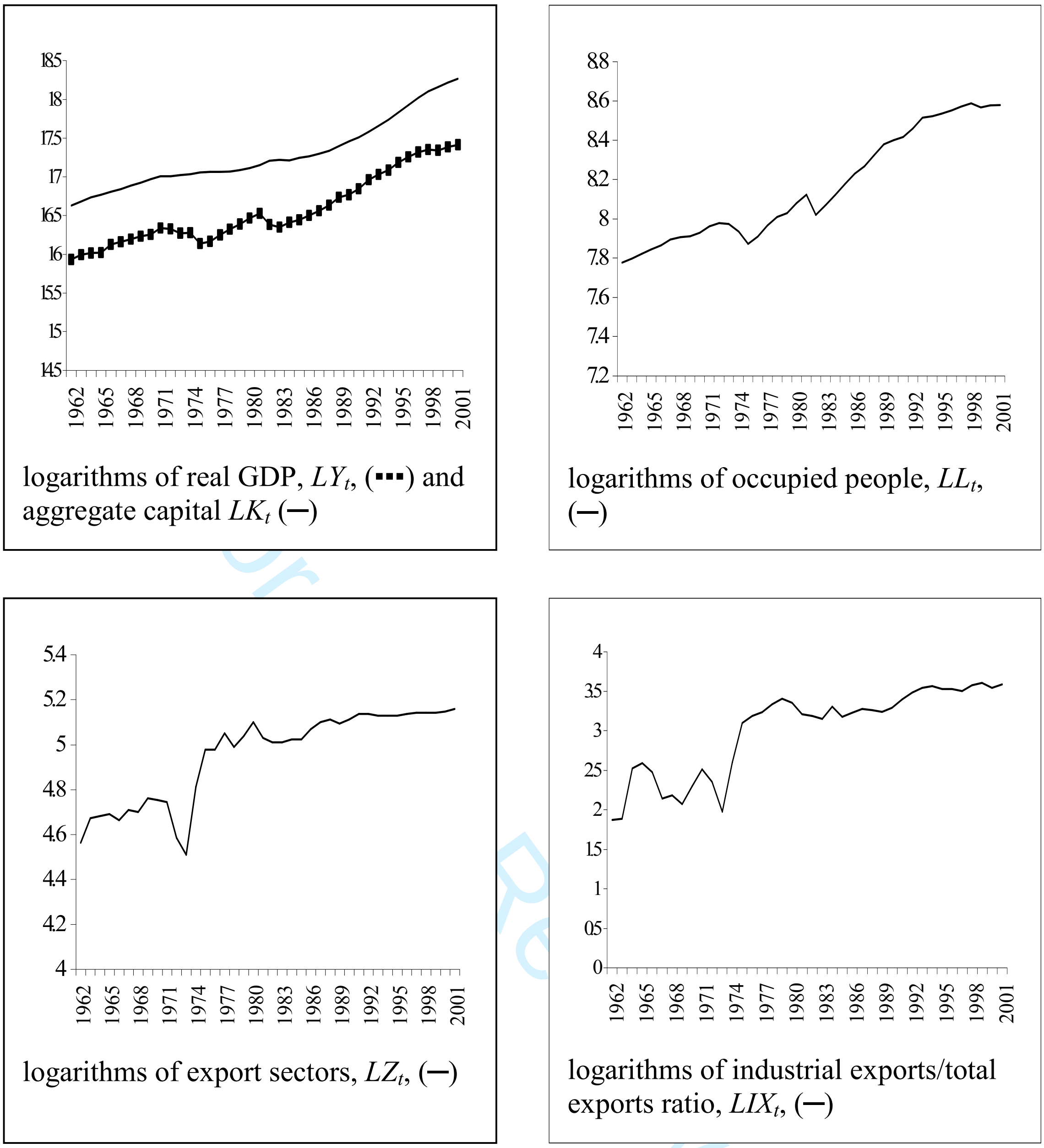


Fig. 2. *Time series used*

Table 1. *Export performance indicators: 1987-2001*

	1987	1988	1989	1990	1995	2000	2001
Number of exported products	1440	1487	1490	2796	3647	3749	3749
Number of exporting firms	3666	3461	3462	4125	5817	5666	6009

Source. PROCHILE (2003)

Table 2. *Perron (1997) unit root test*

Series	Model	Break Year	Dummy Variables	Test Statistic $t_{\hat{a}}$	Critical Value 5% (1%)	Result
Levels						
LY_t	(8)	1971	*du*72, *d*72	-2.89	-5.23 (-5.92)	$I(1)$
LK_t	(9)	1981	*dt*82	-2.48	-4.83 (-5.45)	$I(1)$
LL_t	(8)	1981	*du*82, *d*82	-3.70	-5.23 (-5.92)	$I(1)$
LZ_t	(8)	1972	*du*73, *d*73	-8.45	-5.23 (-5.92)	$I(0)$ + trend
LIX_t	(8)	1972	*du*73, *d*73	-6.91	-5.23 (-5.92)	$I(0)$ + trend
First Differences						
$\Delta(LY_t)$	(8)	1971	*d*72	-4.45	-3.53 (-4.23)	$I(0)$
$\Delta(LK_t)$	(9*b*)	-	-	-2.64	-1.95 (-2.62)	$I(0)$
$\Delta(LL_t)$	(8)	1981	*d*82	-4.57	-3.53 (-4.23)	$I(0)$

Notes: The dummy variables are specified as follows: *d*72, *d*82, *d*73 are impulse dummy variables with zeros everywhere except for a one in 1972, 1982, 1973. *du*72, *du*82, *du*73 are 1 from 1972, 1982, 1973 onwards and 0 otherwise. *dt*82 is 0 before 1982 and *t* otherwise. Critical values for the levels are provided by Perron (1997). Critical values for the first differences are from MacKinnon (1991). For the first differences only impulse dummy variables were included in the regression. Impulse dummy variables, that is those with no long-run effect, do not affect the distribution of the MacKinnon Test statistics.

Table 3. *Kapetanios (2002) unit root test*

Series	Break Year	Break Year	Dummy Variables	Test Statistic $t_{\hat{a}}$	Critical Value 5% (1%)	Result
Levels						
LY_t	1973	1981	*du*73, *du*82	-3.59	-5.69 (-6.16)	$I(1)$
LK_t	1974	1981	*dt*75, *dt*82	-2.95	-6.11 (-6.59)	$I(1)$
LL_t	1973	1981	*du*74, *du*82	-2.73	-5.69 (-6.16)	$I(1)$
LZ_t	1971	1973	*du*72, *du*74	-11.9	-5.69 (-6.16)	$I(0)$ + trend
LIX_t	1971	1973	*du*72, *du*74	-6.45	-5.69 (-6.16)	$I(0)$ + trend
First Differences						
$\Delta(LY_t)$	1973	1981	*d*74, *d*82	-3.85	-3.53 (-4.23)	$I(0)$
$\Delta(LK_t)$	1974	1981	*d*75, *d*82	-3.54	-3.53 (-4.23)	$I(0)$
$\Delta(LL_t)$	1973	1981	*d*74, *d*82	-4.90	-3.53 (-4.23)	$I(0)$

Notes: The dummy variables are specified as follows: *d*74, *d*75, *d*82, are impulse dummy variables with zeros everywhere except for a one in 1974, 1975, 1982. *du*72, *du*73, *du*74, *du*75, *du*82 are 1 from 1972, 1973, 1974, 1975, 1982 onwards and 0 otherwise. *dt*82 (*dt*75) is 0 before 1982 (1975) and *t* otherwise. Critical for the levels are provided by Kapetanios (2002). Critical values for the first differences are from MacKinnon (1991). For the first differences only impulse dummy variables were included in the regression. Impulse dummy variables, that is those with no long-run effect, do not affect the distribution of the MacKinnon Test statistics.

Table 4. *Johansen's trace-test for multiple cointegrating vectors; variables: LY_t, LK_t, and LL_t*

Statistics	Critical Value 95% (99%)	Null Hypothesis	Alternative Hypothesis
36.173***	29.68 (35.65)	$r = 0$	$r \geq 1$
14.770	15.41 (20.04)	$r \leq 1$	$r \geq 2$

Notes: *** indicate a rejection at the 99% critical value. Critical values are taken from Osterwald-Lenum (1992).

Table 5. Johansen's trace-test for multiple cointegrating vectors; variables: lix_t, lz_t

Statistics	Critical Value 95% (99%)	Null Hypothesis	Alternative Hypothesis
29.939***	15.41 (20.04)	$r = 0$	$r \geq 1$
9.755***	3.76 (6.65)	$r \leq 1$	$r \geq 2$

Notes: *** indicate a rejection at the 99% critical value. Critical values are taken from Osterwald-Lenum (1992).

Table 6. *DOLS procedure results*

α	β	γ	δ
0.75^{***} (10.16)	0.45^{***} (3.61)	0.49^{**} (2.68)	0.08^{**} (2.10)

$\bar{R}^2 = 0.99$ $SE = 0.026$ $DW = 1.77$ $JB = 0.27\ (0.87)$

$ARCH(1) = 2.84\ (0.11)$ $ARCH(2) = 1.87\ (0.17)$ $ARCH(4) = 0.17\ (0.95)$

$LM(1) = 0.36\ (0.55)$ $LM(3) = 1.19\ (0.34)$ $White = 1.04\ (0.53)$